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USING QUIT RATES TO SET COMPENSATION LEVELS IN THE PUBLIC SECTOR

Kathleen C. Utgoff

CNA
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The government can not base the amount it pays its employees on the market value of the services it provides. Instead, it tries to set pay for each job at a level equal to pay for comparable work in private employment. Although a survey of pay for private sector workers is taken each year, there are indications that the survey is faulty and that government pay levels are too high. One indication is that quit rates are considerably lower in the government than in the private sector. Quit rates are a function of pay levels. Other things equal, comparable quit rates indicate comparable pay.

20 The evidence presented in this paper shows that one reason for the lower quit rates in the government is that the government is such a large employer. Large firms have lower quit rates, probably because transfer and promotion are easy within a large firm. Even when firm size is taken into account, however, the quit rate in the government remains lower than the quit rate for equivalent private sector workers, suggesting that compensation (pay plus other benefits) for government is above private compensation. The differential--shown in this paper to be 15 percent--is in line with findings in other studies that make direct pay comparisons, but these results must be considered tentative until other factors that influence the quit rate have been included in the analysis.

The estimating equations in this study are derived from a simple model of labor market behavior. Some workers accept low pay while searching for higher paying jobs, and firms with low turnover costs offer low pay and have high quit rates. This model makes it clear that efficient behavior for the government does not imply either the same pay or total compensation as the private sector. The optimal pay level minimizes the sum of turnover and compensation costs.

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USING QUIT RATES TO SET COMPENSATION LEVELS IN THE PUBLIC SECTOR

Kathleen C. Utgoff

Prepared under contract N00014-79-C-0448 for:

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INTRODUCTION[1]

By law, federal pay is supposed to be comparable to pay in the private sector. Most people believe, however, that except for the highest grade levels, government workers are paid more than their civilian sector counterparts. This belief is supported by studies that show government jobs pay more than equivalent private jobs, and also, that government workers are paid more than private workers with the same skills.[2]

Federal pay is set by a complex and controversial process. White-collar, blue-collar, postal, and military employees are each covered by a different pay system. The most important of these systems, in terms of its budgetary impact, is the pay system for setting the salary of white-collar employees known as the General Schedule. The General Schedule applies to over half the civilian federal workers, or 1.4 million white-collar employees. In addition, the pay of the two million uniformed military is also linked to adjustments in the General Schedule.[3] General Schedule payroll costs exceed $30 billion per year.[4]

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[1]This paper could not have been written without the help of my colleagues at CNA. Dick Thaler, a consultant to CNA, had a large impact on the design and execution of the paper. Ed Berger and Marie Makurath worked on the project as research assistants. Frederick Miller, Richard Hunter, and George Borjas also contributed. I am grateful to all of them.

[2]See Fogel and Levin [6], Perloff [14], Quinn [15], and Smith [17].

[3]The Federal Wage (blue-collar) is based on local pay rates, and the Postal Wage is determined by collective bargaining.

[4]Statistics and procedures for federal pay systems can be found in CBO Background Paper No. 19, "The Federal Pay System: Adjustment Procedures and Impacts of Proposed Changes," [5].

The General Schedule pay rates are adjusted annually according to procedures established by the Federal Pay Comparability Act of 1970. Each year a BLS survey is taken to determine pay rates in the private sector. This survey is used to establish a "pay line" that is supposed to result in federal pay that is comparable to private pay and, in addition, maintain appropriate grade distinctions. Both the survey design and its use to calculate the pay line have been criticized as being biased in favor of higher pay levels.[1]

It's hard to imagine that there will ever be agreement on the comparability of federal and civilian compensation, especially if comparability is defined to include nonpecuniary benefits and fringe benefits in the compensation package. One suggestion for circumventing the disputed pay-setting process is to use the quit rate as a measure of comparability. It is clear, on the basis of previous economic models and empirical studies, that quits are a function of workers' perceptions about relative pay--the higher the relative pay, the lower the quit rate.[2] Thus, it seems logical to use the quit rate to see if workers consider jobs in the private and public sector comparable.

This paper presents additional evidence that quit rates are a function of relative pay. The elasticity of the quit rate with respect to pay levels is estimated using aggregate and individual data, both across industries and within a single industry (steel). The negative relationship between quit rates and pay also holds within the government--agencies with salary premiums also have low quit rates.

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[1] See [5] and [6] for some of these criticisms.

[2] See Parsons [13] for a review of studies of the economic determinants of the quit rate.

This paper compares the quit rate in the government with both the BLS quit rate in manufacturing and a measure of the quit rate in the steel industry. It is clear that the government quit rate is considerably below private sector quit rates. What is not clear is how much of the difference is due to higher compensation levels, and how much is due to other differences between the public and private sector that also influence quit rates. One obvious difference between the government and the private sector is the size of the government compared to the size of the average private sector firm. Since the probability of intrafirm transfer is greater in a large firm, the quit rate is mechanically related to the number of positions within a firm, or firm size. According to the estimates presented here, much of the difference between the quit rate in the government and the private sector can be accounted for by the larger size of the government.

The empirical estimates in this paper are based on differences in quit rates across industries or firms. Some employers choose high quit rates and low pay levels while others choose low turnover and high pay levels. This paper contains a simple model that explains how firms choose among the set of feasible wage-quit combinations. A long-observed empirical regularity is that large firms appear to pay high wages. Why then, if the government is supposed to mimic the profit-induced efficiencies of the private sector, shouldn't the government pay high wages? Answering this question is a major objective of this paper.

LABOR FLOWS IN THE GOVERNMENT AND THE PRIVATE SECTOR

The turnover statistics shown in table 1 support the contention that government jobs are more attractive than private jobs. Turnover in the government is much lower and less cyclically sensitive than turnover in manufacturing industries, the only private sector where turnover statistics are routinely collected. Turnover statistics are presented for two years to show the cyclical sensitivity of labor flows in the private sector. Flows--both separations and accessions--average about fifty percent of the employment stock in a private firm. All labor flows in the private sector are about five times higher than in the government. Quits, the proposed alternate measure of pay comparability, are always considerably higher in the private sector than in government, but the relationship is extremely volatile over the cycle. Quits are low during recessions both because of reduced job opportunities and because young workers, who are most likely to quit, are already out of work.[1]

There is an important distinction, however, between the private and government statistics reported in table 1: the government can be considered a single large firm, but the manufacturing turnover statistics are for many, often small, firms. The quit rate is mechanically related to firm size because of the increased possibility of intrafirm transfer in large firms. If an employee is dissatisfied or wants to move up the management ladder, alternate opportunities are more likely in a large firm. Thus, it is possible that at least

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[1]For an estimate of the relative importance of the two causes of cycles in quits, see Jacobson [10].

some of the difference between quit rates in the government and private industry is just due to the fact that the government is a much larger employing unit.[1]

TABLE 1

LABOR STOCKS AND FLOWS[a]

	Government (GS - permanent) 1974	Government (GS - permanent) 1975	Private (Total manufacturing) 1974	Private (Total manufacturing) 1975
	(%)		(%)	
Total separations	11	10	58.8	50.4
Quits	4	4	28.8	16.8
Layoffs	.03	.03	18	25.2
Other	6.97	5.97	12	8.4
Total accessions	12	9	50.4	44.4
New hires	7	5	38.4	24.0
Rehires	3	3	12	20.4
Other	2	1	--	--
Employment stock	1.28M	1.3M	20.1M	18.3M

[a]In both data sources, the stocks and flows are inconsistent--the employment at t_0 plus the sum of the flows for N years is less than the employment stock at t_N. If stock figures are accurate, this means that recorded accessions are too low or separations are too high. For a discussion of this problem in the BLS data, see Brechling [3].

Sources: Total Manufacturing from Employment and Earnings. Government data from semi-annual turnover statistics released by OPM.

[1]Quits must ultimately be related to firm size since only labor force dropouts would be counted as quits if everyone worked for a single firm.

There is not a great deal of data on firm size that can be linked to quit propensity. What there is, however, indicates that firm size is related to quits. Table 2 reports correlations from two different sources of firm size data: (i) Census of Manufacturing survey results matched with BLS industry quit rates at both the 3- and 4-digit levels and (ii) state data on firm size from Unemployment Insurance records matched with BLS total manufacturing quit rates. In these data sets, variation in firm size comes from (i) differences across industries in firm size and (ii) from differences across states in average firm size.

In both sets of data there is a strong negative relationship between firm size and quit rates, but the other two correlations are very different for the two types of data. The cross-industry correlations show the expected pattern; the correlation of both quits with earnings and firm size with earnings have the signs found in most other studies; but the cross-state correlations do not conform to expectations. Most previous studies have shown a strong negative relationship between quit rates and earnings.[1] In addition, a positive association between firm size and earnings has often been noted.[2]

---

[1]See Brechling and Jacobson [2] and Parsons [13].

[2]See, for example, Lester [11]. I know of two in-process dissertations on the relationship of firm size and earnings: Miller [12] claims that a pure size differential exists for only white-collar workers; he believes that blue-collar workers earn more in large firms because large firms hire better blue-collar workers. Garen [7], on the other hand, finds that large firms do pay a premium. In his model, large firms have higher screening costs because they have control problems.

TABLE 2

CORRELATIONS
QUIT RATES, FIRM SIZE, EARNINGS

I. CROSS INDUSTRY (1972)

QUIT: BLS-average monthly quit rates, 1972
SIZE: 1972 Census of Manufactures--number of employees divided by number of establishments
EARN: BLS-average hourly earnings, 1972

70 4-Digit Industries

	QUIT	SIZE	EARN
QUIT	1.00		
SIZE	-.36	1.00	
EARN	-.83	.32	1.00

100 3-Digit Industries

	QUIT	SIZE	EARN
QUIT	1.00		
SIZE	-.22	1.00	
EARN	-.84	.19	1.00

II. CROSS STATE (1974, N = 42)

QUIT: Unpublished BLS monthly quit rates by state, 1974 (not available for all states)

SIZE: Unpublished data from administrative unemployment insurance forms; total employment divided by number of reporting units--first quarter 1975

EARN: BLS-straight-time average hourly earnings, 1974

	QUIT	SIZE	EARN
QUIT	1.00		
SIZE	-.43	1.00	
EARN	-.02	-.11	1.00

See appendix A for a more complete description of data definitions and sources.

These unusual results can be explained with a simple model of worker and firm behavior, presented below. This model of the equilibrium relationships among quit rates, firm size, and earnings is used as a basis for the empirical work that follows and for the derivation of policy implications from the empirical results.

THE EQUILIBRIUM DETERMINATION OF QUITS AND EARNINGS BY FIRM SIZE

The motivation for this study was the belief that quit rates are related to relative pay--the more a firm (or the government) pays relative to its competition, the lower its quit rate will be. For this to be an equilibrium relationship, and not just a temporary increase in quits following a shift in relative prices, some firms must deliberately choose higher quit rates and, therefore, turnover costs. Moreover, some workers must be willing to accept jobs in firms which choose low wages and high quit rates.

Workers will accept a low wage if working is better than unemployment and job search can continue after the worker accepts a job. The higher the current wage, the lower is the probability that the worker will either seek or receive a better offer and quit.[1] Two workers with the same skills can have different quit propensities because of differences in tastes and search costs. The aggregation of quit propensities across individuals yields a relationship between quits and wages for a firm. Firms that choose low wages will have high turnover, firms that choose high wages will have low turnover.

---

[1] Ken Burdett [4] has developed a search model with these properties. The model has two switchpoint (reservation) wages: X and Y. The job seeker accepts any wage above X, but he continues to seek a better job at any wage between X and Y. At any wage higher than Y, the worker feels there is little chance of receiving a better offer, so he stops looking and rarely quits.

We will show how a firm makes this choice by modeling the behavior of a firm that has already determined the desired size of its operation, taken here to mean a desired number of employee slots. Over any given time period, increases in turnover will increase the number of individuals ever-on-the-payroll per slot. To maximize profits, the employer will minimize costs (C) per slot, where

$$C = W + TC(q) \tag{1}$$

and

W = the wage rate per period, where W includes the cost of providing fringe benefits

TC = turnover costs, the fixed costs of hiring plus lost productivity due to training time[1]

q = the quit rate per period.

The quit rate (q) is a negative function of W and, by hypothesis, a negative function of firm size:

$$q = q(W,FS),\ \delta q/\delta W < 0,\ \delta q/\delta FS < 0\ . \tag{2}$$

The employer chooses an optimum $W = W^*$ to minimize C. This cost minimization implies that

$$\delta C/\delta W = 1 + TC(\delta q/\delta W) = 0$$

---

[1]Equation (1) applies only to firms where turnover costs are fixed (independent of the level of q). In firms where marginal productivity depends on tenure (and, hence, q), the profit maximizing problem is more complex. The major points in this paper can be made without introducing this complexity.

or

$$TC = -\delta W^{*}/\delta q \ . \tag{3}$$

Figure 1 illustrates the relationship between turnover costs, wages, and firm size that results from this optimizing procedure. The curves FS_1, FS_2, and FS_3 represent the choice set of wages and quit rates for three different size firms. Small firms must choose along FS_1; FS_2 and FS_3 are for larger firms.[1] The straight lines in the figure represent combinations of wages and hiring costs that produce the same total cost per employment slot (C). Each straight line is the set of qs and Ws that satisfy equation 1 for a given C (which is equal to the intercept along the W axes). The slope of the line is determined by turnover costs (TC); an increase in turnover costs flattens the cost line.

The problem for a firm of given size $(\overline{FS})$ and given turnover costs $(\overline{TC})$ is to choose a point along the $q(w,\overline{FS})$ curve. The point that will minimize C is the point where the set of parallel lines determined by $TC = \overline{TC}$ is tangent to the $q(W, \overline{FS})$ curve. Points A and B represent optimum points chosen by two different firms of equal size (FS_3) but with different turnover costs. The point B would be chosen by a firm with high turnover costs; points A would be chosen by a firm with low turnover costs.[2]

---

[1]As discussed above, greater possibility for transfer in large firms allows them to obtain lower turnover for any wage scale.

[2]CNA is a good example of a firm with relatively high turnover costs since a security clearance must be obtained for each employee.

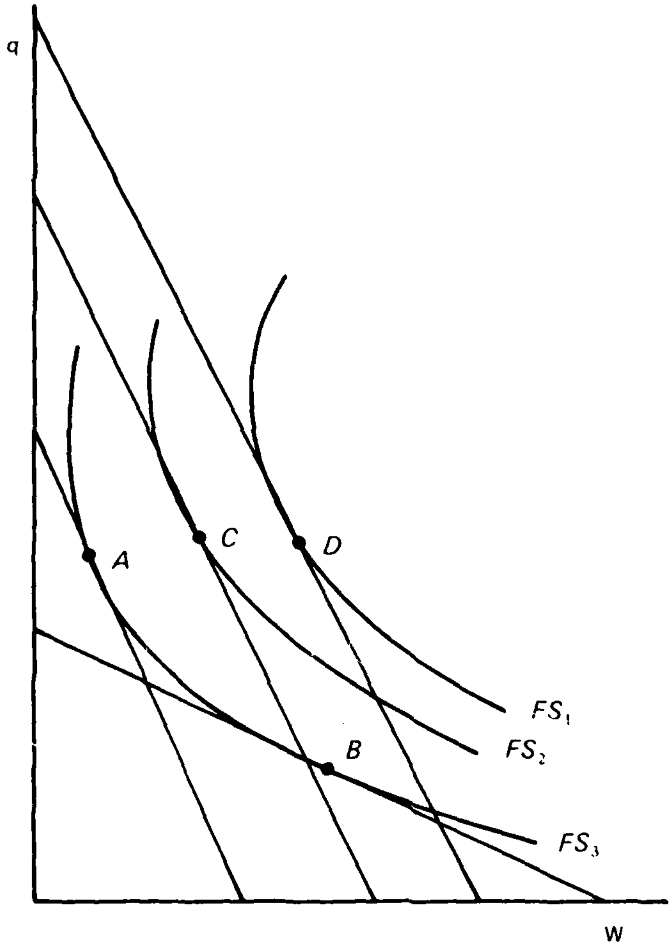

FIG. 1:

The points A, C, and D represent the set of points chosen by the three different size firms with the same turnover costs. Point D is chosen by a small firm, point A by a large firm.[1]

[1]Although simple, this model has a number of appealing properties. The world it describes has a nondegenerate, equilibrium distribution of wages, a crucial, yet unexplained feature of the world described by search theory models. In addition, quits can be explained without resorting to unfulfilled expectations or mistakes on the part of workers; they are systematic rather than totally random. The model can also be used as a framework for dual labor market theory, since it describes a market where some firms, particularly those with little specific capital accumulation, pay low wages and have high turnover.

EMPIRICAL IMPLICATIONS OF THE OPTIMIZATION MODEL

The simple optimizing procedure illustrated in figure 1 has been explained in detail because it provides an explanation for the seemingly diverse correlations in table 2. As shown in figure 1, the observed relationship between quits and wages will depend on the variance of turnover costs relative to firm size in the data. If firm sizes are relatively constant in the data and turnover costs vary, the data will look like points A and B, and there will be a negative relationship between quits and earnings. If on the other hand, turnover costs are relatively constant, the data points will look like points A, C, and D, and the relationship between quits and earnings will be much flatter; it could even be positive.[1] Thus, the diverse results in table 2 could be explained by the fact that the cross-industry data trace out points like A and B, producing a strong negative relationship between earnings and quits; while the cross-state data traces out points like A, C, and D, producing no relationship between earnings and quits. This explanation would imply that the variance in turnover costs relative to the variance in firm size is greater across industries than states. Since states are probably more homogeneous, more like small labor markets than industries, this is not a particularly strained explanation.

Figure 1 illustrates the classic identification problem when two variables like q and W are endogenous. They are simultaneously determined by both equations (1) and (2). To

[1]Note that the empirically observed positive relationship between firm size and earnings is explicable with this model only if we assume that turnover costs are positively related to firm size.

judge pay comparability on the basis of quit rates, we need an estimate of the pure effect of firm size on quit rates, or the distance between points like D and B in figure 1. This means that equation (2) has to be identified. Single equation, OLS estimates of the quit rate as a function of wages (or earnings) and firm size, like those in table 3, will not, in general, identify equation (2). All three equations imply that quit rates are a negative function of earnings and firm size, since all the coefficients are negative and significant.[1] These estimates may be biased, however; if they are, the estimated coefficient of firm size is biased away from zero (the bias is negative).[2] Given the previous explanation of the correlations in table 2, it is not surprising that the size coefficient is smaller (more negative) for the cross-state data. The low simple correlation between quits and earnings in the raw cross-state data means that estimates using these data are less likely to identify equation (2) than estimates from the cross-industry data.

[1]Whenever possible all equations were estimated in ln-ln form. These equations allow for declining marginal effect of firms size and earnings; they also allow for a comparison of equations estimated in different units since the coefficients are elasticities.

[2]See appendix C for a proof.

TABLE 3

OLS REGRESSIONS[a]

ESTIMATED EQUATION $\ln(QUIT) = C_0 + C_1 \ln(EARN) + C_2 \ln(SIZE)$

Coefficients

Variable	4-DIGIT	3-DIGIT	CROSS-STATE
EARN	-2.048 (12.62)	-2.504 (16.63)	-1.179 (3.17)
SIZE	-.178 (4.94)	-.173 (5.31)	-.508 (4.01)
CONSTANT	15.796	18.533	12.411
$\bar{R}^2$	.759	.785	.324
N	70	100	42

[a]See table 2 and appendix A for variable definitions. The numbers in parentheses below the coefficients are the absolute values of the t-statistics.

SIMULTANEOUS EQUATION ESTIMATES

Simultaneous equation techniques are required to obtain an unbiased estimate of the effect of firm size on quit rates. Since there is not enough additional exogenous data to identify equation (2) using the BLS data, we have used two samples that record a great deal of information on individual workers. The simultaneous equation estimates from these data are reported below.

ESTIMATES FROM THE SOCIAL SECURITY DATA

The LEED (Longitudinal Employer-Employee Data) file is a one percent sample of Social Security records. This sample has been used extensively by Jacobson [9] to study workers who leave (either a quit or layoff) a number of different industries. This sample is so large that Jacobson was able to estimate firm size by counting up the number of individuals with the same firm identification number. The data indicates when an individual separates from a firm but not whether he was laid off or quit. To identify quits, Jacobson calculated the yearly change in employment in each worker's firm. Any person who left a firm with rising employment can be identified as a quit, since firms with rising employment very rarely have layoffs.[1]

[1] A few of those who separate from firms with rising employment are discharges. The theory described above applies to discharges as well as quits since both generate turnover costs. In practice, discharges are rare in steel and nonexistent in the government. See "Firing a Federal Employee: The Impossible Dream" by Reed [16].

The results presented below are based on observation of individuals in steel (SIC 3312) firms with rising employment; observations on individuals from firms with steady or falling employment were omitted from the sample whether they changed firms or not.[1] The sample was further restricted to observations from SMSAs that could be matched to BLS labor market data. The final sample consisted of 4,658 observations from 1963 to 1970.[2]

Table 4 presents the coefficients from a two-stage discriminant estimate of equation (4):[3]

$$\ln\left(\frac{Q(L)}{1-Q(L)}\right) = \alpha_0 + \alpha_1 \widehat{\ln EARN(L)} + \alpha_2 \ln SIZE(L) \qquad (4)$$

where $Q(L)$ is the probability that the individual quits and $\widehat{\ln EARN(L)}$ is the estimated earnings from the first stage, an OLS regression that estimates earnings as a function of firm size, age, race, tenure, the stability of previous employment, some labor market variables, and seven year dummies.

Again, the estimated coefficients support the hypothesis that quits are a negative function of earnings and firm size; both coefficients are negative and significant. The estimated

---

[1] An observation or data point is a person-year. Individuals appear in the data matrix several times, once for each year that they work in a firm with rising employment.

[2] Data from 1957 to 1972 was used to calculate variables (e.g., tenure); the quit observations are for the years cited.

[3] The LEED estimates were obtained using the discriminant option of RAND's ECON package. The discriminant estimates will approximate the parameters estimated by a conditional maximum likelihood logit estimate where the dependent variable is binary. For a discussion of the discriminant function, see Halperin, Blackwelder, and Verter [8].

TABLE 4

LEED REGRESSIONS[a]
(two-stage discriminant)

FIRST STAGE

$$
\begin{aligned}
\widehat{\ln EARN(L)} = -.0015 &- \underset{(2.931)}{.0222}\, SIZE(L) + \underset{(21.137)}{.0921}\, AGE - \underset{(19.239)}{.001}\, AGE^2 \\
&\underset{(4.856)}{.3387}\, DRACE1 - \underset{(3.555)}{.0764}\, DRACE2 + \underset{(10.038)}{.3483}\, DTEN2 \\
&+ \underset{(10.099)}{.5188}\, DTEN3 + \underset{(6.829)}{.4729}\, DTEN4 + \underset{(7.290)}{.3889}\, DTEN5 \\
&+ \underset{(6.425)}{.3953}\, DTEN6 + \underset{(13.184)}{.3796}\, DTEN7 + \underset{(3.276)}{.0006}\, LQIM1 \\
&- \underset{(4.625)}{.0008}\, LQIM2 + \underset{(3.050)}{.0024}\, GROW1 + \underset{(.2820)}{.0003}\, CITSIZ \\
&+ \underset{(1.841)}{.008}\, CYCY + \underset{(1.0314)}{.0005}\, CYCM1 - \underset{(1.368)}{.0352}\, YD3 - \underset{(.602)}{.0217}\, YD4 \\
&+ \underset{(.958)}{.0456}\, YD5 + \underset{(2.560)}{.1164}\, YD6 + \underset{(2.617)}{.1308}\, YD7 + \underset{(4.254)}{.1763}\, YD8 \\
&+ \underset{(2.144)}{.1286}\, YD9; \quad \bar{R}^2 = .3872
\end{aligned}
$$

SECOND STAGE

$$
\ln\left(\frac{Q(L)}{1-Q(L)}\right) = \underset{(15.319)}{3.6676} - \underset{(23.853)}{2.2012}\, \widehat{\ln EARN(L)} - \underset{(6.707)}{.2286}\, \ln SIZE(L)
$$

---

[a]See appendix A for variables definitions. The numbers in parentheses below the coefficient are the absolute value of the t-statistics.

elasticities are quite close to the estimated elasticities from the BLS equations (the coefficients of the BLS equations). The elasticities from the LEED equation are equal to the coefficient times $1-Q(L)$ since

$$Q(L) = \frac{e^{\Sigma\alpha_i X_i}}{1+e^{\Sigma\alpha_i X_i}} \tag{5}$$

$$\frac{dQ(L)}{dX_i} = \alpha_i \cdot 1-Q(L) \cdot Q(L)$$

and, thus, at the sample mean of $Q(L)$, which is .1958,

$$\frac{d\ln Q(L)}{d\ln EARN(L)} = \alpha_1 \cdot 1-Q(L) = 1.77$$

$$\frac{d\ln Q(L)}{d\ln SIZE(L)} = \alpha_2 \cdot 1-Q(L) = .1838$$

ESTIMATES FROM THE MICHIGAN SURVEY OF WORKING CONDITIONS

The 1969-1970 University of Michigan Survey of Working Conditions (SWC) was also used to obtain estimates of the effect of firm size on the quit rate. The SWC is a detailed survey of workers undertaken from December 1969 to January 1970. Farmers, self-employed workers and government workers were excluded from the observations used to obtain the estimates. The regression estimates are based on 1045 survey responses. One question in the survey dealt with the workers quit intentions over the next year. The quit variable [Q(M)] is equal to 1 if the worker said he was either very likely or somewhat likely to quit.

Several quit functions were estimated with personal and job characteristics as independent variables. The regression variables were the same as those used by Viscusi [21] to estimate the effect of job hazards on quit propensities.

Table 5 gives two-stage and single-equation estimates of a logistic probability function where Q(M) is described by the same function as Q(L) in equation (5). Many of the variables have the expected effect on quit propensities; workers are less likely to say that they intend to quit if they have high tenure, if they feel their jobs are secure, or if they are older. Quit intentions are stronger if a job is physically arduous, if it is dangerous, or if the worker has a health problem. In contrast to the other empirical results presented here, there is no evidence from the SWC that firm size affects quit rates.[1]

It is possible, however, that Q(M) reflects workers dissatisfaction about their current job and that in big firms this dissatisfaction will be alleviated by a transfer within the firm, a transfer that is impossible within a small firm. This means that actual quits could be related to firm size even though quit intentions are not.

---

[1] Q(M) was also estimated using OLS with the same independent variables. All equations were estimated with and without the tenure variables since tenure is a function of the quit rate. In addition, many equations were estimated with both a continuous size variable (equal to the mid-point of the interval class or the lower bound of the highest class) and with interval dummies. There was no evidence that size affected quit propensities for the total sample. Since Viscusi's work ([20], [21]) indicates a difference between blue- and white-collar workers, separate equations were estimated for each group. A small effect of firm size on quit rates was observed for blue-collar workers (about half the sample). Most general schedule employees would, however, be considered white-collar workers.

TABLE 5

SWC REGRESSIONS
(TWO-STAGE LOGISTIC ESTIMATES, N = 1045)

FIRST STAGE

$$
\begin{aligned}
\widehat{\ln \mathrm{EARN}(M)} = 8.40 &- \underset{(4.44)}{.267}\ \mathrm{SIZE1} - \underset{(2.10)}{.116}\ \mathrm{SIZE2} - \underset{(1.57)}{.100}\ \mathrm{SIZE3} \\
&- \underset{(1.46)}{.080}\ \mathrm{SIZE4} - \underset{(.20)}{.013}\ \mathrm{SIZE5} - \underset{(1.04)}{.075}\ \mathrm{SIZE6} \\
&+ \underset{(7.14)}{.049}\ \mathrm{AGE} - \underset{(7.20)}{.0006}\ \mathrm{AGE2} - \underset{(2.76)}{.128}\ \mathrm{RACE} - \underset{(1.26)}{.181}\ \mathrm{TEN1} \\
&- \underset{(1.37)}{.105}\ \mathrm{TEN2} - \underset{(3.02)}{.191}\ \mathrm{TEN3} - \underset{(3.09)}{.184}\ \mathrm{TEN4} - \underset{(2.17)}{.135}\ \mathrm{TEN5} \\
&- \underset{(2.52)}{.148}\ \mathrm{TEN6} - \underset{(2.92)}{.160}\ \mathrm{TEN7} - \underset{(2.5)}{.089}\ \mathrm{BCOL} - \underset{(.088)}{.026}\ \mathrm{PHYSC} \\
&+ \underset{(3.93)}{.196}\ \mathrm{DEC1} + \underset{(1.55)}{.076}\ \mathrm{DEC2} + \underset{(.26)}{.014}\ \mathrm{DEC3} + \underset{(.132)}{.056}\ \mathrm{CREAT1} \\
&+ \underset{(1.43)}{.057}\ \mathrm{CREAT2} + \underset{(.99)}{.043}\ \mathrm{CREAT3} + \underset{(.99)}{.030}\ \mathrm{DANGER} \\
&+ \underset{(2.44)}{.084}\ \mathrm{FRINGE} + \underset{(.86)}{.036}\ \mathrm{FAST} + \underset{(.275)}{.079}\ \mathrm{SECUR} \\
&+ \underset{(2.26)}{.069}\ \mathrm{TRAIN} + \underset{(3.71)}{.122}\ \mathrm{UNION} - \underset{(1.48)}{.075}\ \mathrm{HEALTH} \\
&- \underset{(2.91)}{.139}\ \mathrm{MARITAL} - \underset{(9.17)}{.761}\ \mathrm{ED1} - \underset{(7.54)}{.602}\ \mathrm{ED2} - \underset{(7.04)}{.489}\ \mathrm{ED3} \\
&- \underset{(5.54)}{.349}\ \mathrm{ED4} - \underset{(4.25)}{.282}\ \mathrm{ED5} - \underset{(.100)}{.072}\ \mathrm{ED6} \\
&- \underset{(19.16)}{.592}\ \mathrm{SEX}; \quad \bar{R}^2 = .556
\end{aligned}
$$

SECOND STAGE

$$
\ln\left(\frac{Q(M)}{1-Q(M)}\right) = 3.35 - \underset{(3.21)}{.468}\ \widehat{\ln \mathrm{EARN}(M)} - \underset{(.820)}{.029}\ \ln \mathrm{SIZE}(M)
$$

TABLE 5 (Cont'd)

SINGLE EQUATION LOGISTIC ESTIMATES

$$\ln\left(\frac{Q(M)}{1-Q(M)}\right) = \underset{}{.924} - \underset{(.47)}{.021\ \text{SIZE}} - \underset{(6.49)}{.041\ \text{AGE}} + \underset{(.238)}{.533\ \text{RACE}}$$

$$- \underset{(1.78)}{.152\ \text{BCOL}} + \underset{(2.05)}{.315\ \text{PHYSC}} + \underset{(1.41)}{.081\ \text{NODEC}}$$

$$+ \underset{(.72)}{.051\ \text{CREAT}} + \underset{(2.39)}{.369\ \text{DANGER}} - \underset{(2.20)}{.386\ \text{FRINGE}}$$

$$+ \underset{(.89)}{.189\ \text{FAST}} - \underset{(5.06)}{.736\ \text{SECUR}} - \underset{(2.46)}{.396\ \text{TRAIN}}$$

$$+ \underset{(.00)}{.006\ \text{UNION}} + \underset{(2.49)}{.644\ \text{HEALTH}} + \underset{(1.34)}{.293\ \text{MARITAL}}$$

$$+ \underset{(1.19)}{.007\ \text{EDUC}} - \underset{(1.38)}{.246\ \text{SEX}}$$

See appendix A for variable definitions. The numbers in parentheses below the coefficients are the absolute values of the t-statistics.

IMPLICATIONS OF THE EMPIRICAL ESTIMATES

Three of the five estimates of the elasticity of quit rates with respect to firm size are slightly under .20. Although it is dangerous to extrapolate outside the range of the observations, we can use the LEED estimates (which are based on the best data and estimating techniques) to calculate what the quit rate would be in the steel industry if the average firm were as big as the government and everything else were held constant. Using the second stage equation listed in table 3, set SIZE(L) equal to the size of the government (1.3M) and lnEARN(L) equal to the LEED sample mean to obtain

$$\ln\left(\frac{Q}{1-Q}\right) = -2.848 \quad \text{or}$$

$$q = .055$$

This predicted quit rate is about 25 percent nigher than the actual quit rate listed in table 1. The predicted quit rate would be equal to the actual government quit rate (.04) if the earnings variable in the LEED equation were increased from the mean by about 15 percent since

$$\frac{\partial \ln Q}{\partial \ln EARN} = 1.7$$

implies

$$\partial \ln EARN = .147$$

for $\partial \ln Q = .25$.

Thus, if we make the strong assumption that the only difference between the quit rate in steel firms and the government is due to size and compensation, we can infer that the compensation of government workers (pay and other benefits) is about 15 percent higher than the compensation of similar workers in the steel industry. This premium is within the

range of government pay differentials estimated by other researchers using very different techniques.[1]

Another implication of the analysis presented here is that pay comparability or compensation comparability is not an efficient pay-setting principle. There are two aspects of an efficient compensation-setting process: First, the compensation package, for any given level of quits, is made up of a package and pay and fringe benefits that minimizes total costs. This is not likely to be the same mix of pay and fringes for the government and the private sector. The government has a natural advantage in producing some fringes, such as job security. More frequent changes in demand make it relatively difficult for the private sector to provide job security. The second aspect of efficient compensation setting is that the choice among the least-cost packages should be the one that minimizes total costs by balancing turnover and compensation costs. As illustrated in figure 2, the most efficient package of wages and fringes will not in general be an equal or "comparable" package. The private sector is represented by the point P with an annual quit rate of .20 (the mean of Q(L) and close to the BLS average for manufacturing). Our LEED estimates imply that if the government compensation was comparable, it would be at point C. The actual quit rate in the government is .04 (point $\hat{G}$), implying

[1]See footnote 2 on page 1. These studies calculated pay differentials. The calculations above ascribe all differences in quit rates not due to firm size to differences in compensation (both pay and fringe benefits). Thus, the results of this study and past studies may not be directly comparable for two reasons: (i) differences in quit rates due to factors other than compensation and firm size and (ii) differences between the public and private sector in breakdown of the compensation package between pay and other benefits. To the extent that other benefits (such as job security and pension rights) are higher in the government, the implied pay differential is less than 15 percent.

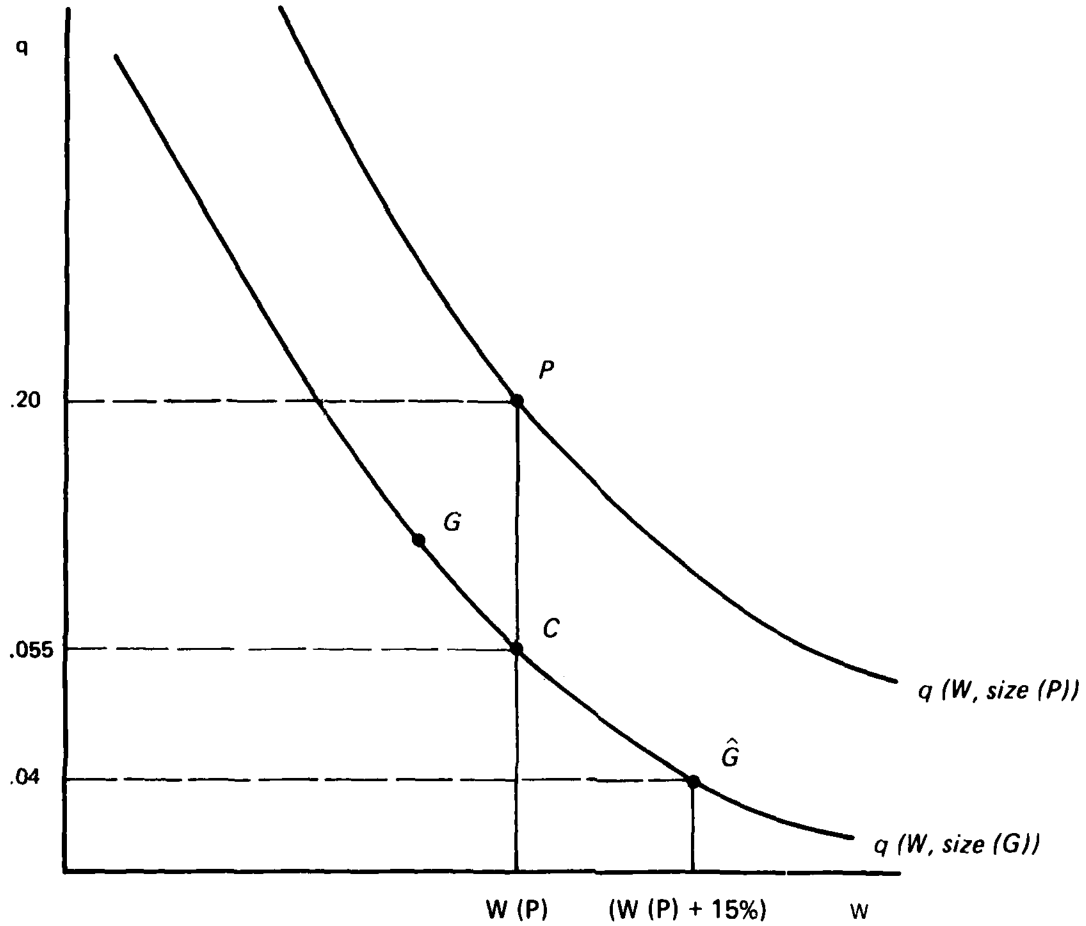

FIG. 2:

more compensation (by about 15 percent). If the government has the same turnover costs as the private sector, the optimum government compensation/quit combination is a point like G, with a quit rate between .055 and .2. If government turnover costs are higher than private costs, the optimum combination is below G.

Turnover costs may be higher in the government because elaborate hiring procedures have been set up to prevent politicians from packing the federal payrolls with friends and relatives.[1] It's unlikely however that turnover costs are high enough to justify a point like $\hat{G}$. Equation (3) and the LEED estimates can be used to obtain estimates of the implied optimal turnover costs at $\hat{G}$ and at points like G and P (with equal turnover costs by assumption).

$$TC = -\delta W^*/\delta q = TC^*$$

from the LEED estimates

$$\frac{\delta \ln Q}{\delta \ln W} = -2.2(1-q)$$

$$\frac{\delta q}{\delta W} = -2.2(1-q)\ \frac{q}{W}$$

thus for P or G

$$TC^*(P) = \frac{W}{2.2(1-q)q} = \frac{W}{.352}$$

[1]Turnover costs may be partially endogenous (dependent on the compensation level); many applicants will apply for the same slot if wages are very high, forcing managers to evaluate many applications to fill a vacancy. These endogenous costs will be reduced when the compensation package is reduced; they should not be used to justify high levels of compensation.

and for $\hat{G}$

$$TC^*(\hat{G}) = \frac{W+.15W}{2.2(1-q)q} = \frac{1.15(W)}{.084}$$

Thus $TC^*(\hat{G})/TC^*(P) = 4.8$.

This means that the turnover costs for the government would have to be almost 5 times higher in the government than in the private sector (steel) to justify the current quit rate.[1]

To summarize, equal or comparable pay is not an efficient pay-setting principle. First, it does not produce an efficient combination of pay and fringe benefits. Second, it does not produce an efficient combination of compensation and turnover costs. It is very unlikely that if the government were to set pay efficiently, it would choose either the same wage or the same quit rate as observed in the private sector because the government has a naturally lower quit rate due to its large size and naturally different forms of nonpay compensation such as job stability. Efficient pay-setting practices will be based on these natural differences.

[1] Even if turnover costs were five times higher in the government, the level of pay could still be inefficient if there was a cheaper way to produce the same quit rate with a different combination of pay and fringe benefits.

EXTENSIONS

Before the quit rate can be used as a pay-setting measure--for determining either comparable pay or efficient pay--the analysis presented here should be extended in several ways.

OTHER DETERMINANTS OF QUIT RATES

A major difference between the government and the private sector is the size of the government, and size seems to have a substantial effect on quit rates. There may be other differences that also affect quit rates. These additional differences could be either differences in the personal characteristics of government and private workers or differences between the government as a firm and private employers.

Personal characteristics like age and sex affect quit rates, and it may be appropriate to make some adjustment to account for differences between the public and private workforce. The problem is complex, however, because government pay-setting practices may cause differences between workers attracted to the public and private sector. Tenure of workers is an extreme example. High pay reduces quit and, hence, increases the average tenure. This means that it may not be appropriate to adjust for all tenure differences when quit rates are used to judge pay scales. Pay-induced tenure differences must be separated from differences in tenure due to other reasons.

In theory, it's easier to adjust the quit rate for differences in the employment environment, though in practice the data collection might be difficult. Table 6 shows the second stage of the LEED discriminant equation from table 4 estimated with several new variables: average earnings in the

area outside the firm (OEARN), the size of the SMSA (LFORCE), and the fraction of the local labor force employed in the firm (SHARE). All have significant effects on the quit rate. These and other differences between the government and the private sector will have to be taken into account when quit rates in the government and private sector are compared.

TABLE 6

EXPANDED LEED ESTIMATES[a]
(two-stage discriminant)

$$\ln\left(\frac{Q(L)}{1-Q(L)}\right) = \underset{(.4726)}{.2712} - \underset{(22.40)}{2.12}\ \widehat{\ln EARN(L)} - \underset{(6.15)}{.217}\ \ln SIZE(L)$$

$$+ \underset{(.72)}{.000}\ (OEARN) + \underset{(6.42)}{.0002}\ OEARN(-1)$$

$$+ \underset{(3.40)}{.0000}\ LFORCE - \underset{(3.78)}{.003}\ SHARE$$

[a]See appendix A for variable definitions. The numbers in parentheses below the coefficients are the absolute values of the t-statistics.

THE SIZE OF THE GOVERNMENT

The preceding analysis treats the government as a single firm. It may be more appropriate to treat each agency within the government (Labor, NSF, Defense) as a single firm, which would make the government smaller and its predicted, size-adjusted quit rate higher.[1] This would increase the implied pay differential for the government.

[1]It may also be appropriate to treat employees within a geographic area as a firm, again reducing the size of the government.

At issue here is whether a transfer within an agency is the same as a transfer between agencies. The probability of quitting $\big(P(Q)\big)$ given job dissatisfaction is

$$P(Q) = 1-[P(T)+P(R)] \tag{6}$$

where $P(T)$ is the probability of transfer between agencies and $P(R)$ is the probability of reassignment within an agency.

If the government is one big firm, then the bracketed term on the right in equation (6) should be independent of agency size since it represents movement to all slots other than the one currently occupied.[1] If, on the other hand, agencies are more like firms, transferring is not as easy as reassignment, and quit rates will be a negative function of agency size.

Table 7 presents the results of the following OLS regression for 21 government agencies:

$$\ln QUIT(A) = b_0 + b_1 \ln EARN(A) + b_2 \ln SIZE(A) .$$

The quit and earnings data come from a paper by George Borjas [1] that is based on personnel transaction records from the Office of Personnel Management (OPM). The variable lnEARN(A) is Borjas' estimate of the earnings differential (relative to (HEW)) for each agency. This measure is adjusted for difference in the quality of workers in different agencies.[2]

---

[1] Statistics for $P(R)$ and $P(T)$ are not available on a consistent basis by agency.

[2] Borjas believes that the pay differentials are based on the political power of the constituency of each agency.

TABLE 7

QUIT RATE AND AGENCY SIZE IN THE GOVERNMENT

REGRESSION ESTIMATES[a]
(OLS)

$$\ln QUIT(A) = \underset{(2.14)}{.70} - \underset{(2.37)}{.070} \ln SIZE(A) - \underset{(4.83)}{2.77} \ln EARN(A)$$

where QUIT(A) = average monthly turnover per hundred workers between 1961 and 1976. Calculated by G. Borjas from data provided by OPM. The rates appear in table 1 of Borjas [1].

SIZE(A) = paid civilian employment in the Federal Government from the Statistical Abstract 1964-1977.

lnEARN(A) = agency coefficient from logarithmic wage equation where HEW is the base agency. These coefficients appear in table 4 of Borjas [1].

[a]The numbers in parentheses below the coefficients are the absolute values of the t-statistics.

The size coefficient is negative and significant but smaller than the BLS or LEED estimates, meaning that the truth lies somewhere between the poles of considering the whole government a firm and considering each agency a firm. Thus, the size-adjusted, comparable wage, quit rate for the government should be more than the .055 that we obtained using the LEED data.

CONCLUSION

The evidence presented here supports the theory that quit rates can be used to measure the relationship between pay in the public and private sector. Empirical evidence from very different data sources is presented to demonstrate that quits are a negative function of compensation levels. Firms with high pay have low quit rates.

Although quit rates are much lower in the government than in the private sector, not all of the difference can be ascribed to differences in pay. Much of the difference can be explained by differences in firm size. Before quit rates can be used as an accurate measure of compensation, other differences must also be taken into account. In practice, determining the quit rate that would indicate pay comparability may be a difficult and disputed exercise, perhaps only useful as a companion to direct pay comparisons.

The quit rate that does signify pay comparability is not the quit rate that the government should choose. Comparability of pay or total compensation is not an efficient compensation principal. The efficient quit rate is the one that minimizes the sum of turnover costs and compensations costs. This is not likely to be the same quit rate as the private sector since the government has a naturally lower quit rate due to its large size and because its turnover costs are probably higher than turnover costs in the private sector.

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APPENDIX A

DATA DEFINITIONS AND SOURCES

APPENDIX A

DATA DEFINITIONS AND SOURCES

INDUSTRY DATA

QUIT	= average monthly 3- or 4-digit quit rates (x 10) for 1972 taken from the 1978 BLS *Employment and Earnings* or if not available in the 1978 version, from the 1975 BLS *Employment and Earnings*.
SIZE	= number of employees in 3- or 4-digit industries divided by number of establishments from the 1972 Census of Manufacturers.
EARN	= average hourly earnings (x 100) 1972, BLS (see QUIT).

STATE DATA

QUIT	= 1974 monthly quit rate in total manufacturing, data not available for all states, unpublished BLS data reproduced in appendix B.
SIZE	= number of employees in total manufacturing for the first quarter of 1975 divided by number of reporting units; unpublished BLS data taken from unemployment insurance forms submitted by employers. Data provided by John Sullivan, BLS, reproduced in appendix B.
EARN	= average hourly earnings of workers on manufacturing payrolls in 1974, from *Employment and Earnings, States and Areas*, 1939-1974 BLS Bulletin 1370-11.

LEED DATA

QUIT(L)	= 1 if a worker separates from a firm with rising employment in year t, 0 if a worker does not separate from a firm with rising employment.
EARN(L)	= worker's average quarterly earnings in t-1.
SIZE(L)	= employment in worker's firm in year t (x .01).
AGE	= age in years.

DRACE1	= 1 if race is unknown.
DRACE2	= 1 if race is nonwhite.
DTEN(i)	= 1 if $(i-1) < \text{TENURE} \leq i$ (for all TENURE ≤ 6); DTEN 7 = 1 if TENURE ≥ 6; TENURE is years of employment in the steel industry since 1956.
LQIM(j)	= the minimum of quarterly intensity measures in year (t-i). Quarterly intensity in quarter j is actual earnings in quarter j divided by a measure of predicted full-time earnings in quarter j. This variable was developed by Jacobson to measure unemployment with earnings records.
GROW1	= accessions minus separations in t-1 for the worker's firm.
CITSIZ	= change in city population 1955-62 (%).
CYCY	= a cycle variable: the SMSA's unemployment rate in 1962 divided by unemployment rate in year t (x 1000).
CYCM1	= CYCY in year t-1.
OEARN	= average quarterly earnings in manufacturing in year t.
OEARN(-1)	= OEARN in t-1.
LFORCE	= labor force in worker's SMSA (x .01): employment plus unemployment.
SHARE	= percent of total SMSA in steel industry (x 10).
YD(i)	= 1 if year = 1961 + i.

SURVEY OF WORKING CONDITIONS (SWC)

QUIT(M)	= 1 if individual says he is very likely or somewhat likely to quit (VAR 0460 = 1 or 3).
EARN(M)	= annual income (VAR 0286).

SIZE(M) = firm size (VAR 0050). The variable is given in interval classes. It was made continuous by taking the midpoint of the interval classes or the lower bound of the highest class.

SIZE(i) i = the coded number of the interval class for (VAR 0050)

1 = 1-9 employees
2 = 10-49
3 = 50-99
4 = 100-499
5 = 500-999
6 = 1000-1999
7 = 2000-and over.

AGE = age in years (VAR 0496).

SEX = 1 if female (from VAR 0537).

BCOL = 1 if worker is blue-collar (VAR 0542).

PHYSC = 1 if a lot of physical effort required (VAR 0069).

DEC(i) = a dummy variable equal to 1 if the individual answer was coded as i in VAR 0071

1 = a lot of decisions
2 = somewhat
3 = a little
4 = not at all.

CREAT(i) = a dummy variable equal to 1 if the individual's answer was coded as i in (VAR 0072)

1 = a lot (creative)
2 = somewhat
3 = a little
4 = not at all.

DANGER = 1 if worker is exposed to dangerous conditions (VAR 0209).

FRINGE = 1 if covered by a pension plan (VAR 0295).

FAST = 1 if individual is required to work fast (VAR 0446).

SECUR = 1 if individual's job is secure (VAR 0421 = 1).

TRAIN = 1 if training program available (VAR 0296).

UNION = 1 if worker belongs to union (VAR 0319).

HEALTH = 1 if worker has health problems (VAR 0405).

MARITAL = 1 if individual is single (VAR 0488).

ED(i) = 1 if individual's education was (i) in VAR 0497

1 = no school or some grade school
2 = completed grade 8
3 = some high school
4 = high school graduate
5 = some college
6 = college graduate
7 = graduate or professional training.

APPENDIX B

STATE DATA

APPENDIX B

STATE DATA

State	State Code	Quit (x 10)	Size	Earn (x 100)
ME	11	37	46	351
NH	12	32	53	364
VT	13	20	40	378
MA	14	21	56	416
RI	15	34	34	362
CT	16	16	67	442
NY	21	16	40	453
PA	23	16	67	442
OH	31	14	77	512
IN	32	17	89	504
MO	34	12	75	562
WI	35	19	64	481
MN	41	26	55	466
IA	42	27	75	491
MO	43	24	61	437
ND	44	38	31	383
SD	45	31	33	379
NE	46	35	50	406
KA	47	34	59	424
DE	51	12	130	458
MD	42	17	94	462
VA	54	28	82	365
NC	56	40	75	328
SC	57	40	84	332
GA	58	37	56	354
FL	59	45	32	374
KY	61	18	83	430
MI	64	34	67	318
AR	71	54	53	330
OK	73	41	48	397
TX	74	37	61	408
MT	81	20	18	495
ID	82	47	32	439
WY	83	38	18	485
CO	84	33	47	458
AZ	86	27	43	440
UT	87	31	40	392
NV	88	34	22	489
WA	91	18	41	523
OR	92	26	34	502
AK	94	92	17	670
HI	95	12	33	425

APPENDIX C

PROOF THAT FIRM SIZE COEFFICIENT HAS A NEGATIVE BIAS IN AN OLS ESTIMATE OF THE EFFECT OF FIRM SIZE AND EARNINGS ON QUIT RATES

APPENDIX C[1]

PROOF THAT FIRM SIZE COEFFICIENT HAS A NEGATIVE BIAS IN AN OLS ESTIMATE OF THE EFFECT OF FIRM SIZE AND EARNINGS ON QUIT RATES

We have assumed in the paper that quits (q) are a Cobb-Douglas function of wages (W) and firm size (FS), or

$$q = AW^{a}FS^{b}e^{\varepsilon} \tag{C-1}$$

and that total costs per worker (C) are a linear function of wages and quit rates, or

$$C = W + dq \tag{C-2}$$

where d is equal to turnover cost (TC). The cost minimizing first order condition is that

$$dAFS^{b}e^{\varepsilon}aW^{a-1\varepsilon} = -1$$

or (solving for W and taking logs)

$$\ln W = \frac{\ln\left(\frac{-1}{dAa}\right)}{a-1} - \frac{b\ln FS}{a-1} - \frac{\varepsilon}{a-1} .$$

The equation to be estimated is

$$\ln q = \ln A + a\cdot\ln W + b\cdot\ln FS + \varepsilon .$$

The asymptotic bias on coefficient vector (from Goldberger, p. 282)[2] is

$$\text{Bias} = \Sigma_{xx}^{-1}\ \text{plim}\left[\frac{X'\varepsilon}{T}\right] . \tag{C-4}$$

where

---

[1]This proof was provided by Jim Jondrow.

[2]Goldberger, Arthur S., <u>Econometric Theory</u>, New York, John Wiley & Sons, Inc., 1964.

Σ_{xx} = covariance matrix of right-hand size variables (including the column of 1s for the intercept). Subscript 0 refers to intercept, 1 to "W" and 2 to "FS."

$$\Sigma_{xx} = \begin{bmatrix} \sigma_{00} & \sigma_{01} & \sigma_{02} \\ \sigma_{10} & \sigma_{11} & \sigma_{12} \\ \sigma_{20} & \sigma_{21} & \sigma_{22} \end{bmatrix} = \begin{bmatrix} 0 & 0 & 0 \\ 0 & \sigma_{11} & \sigma_{12} \\ 0 & \sigma_{21} & \sigma_{22} \end{bmatrix}$$

$$\text{plim}\left(\frac{X'\varepsilon}{T}\right) = \text{vector of asymptotic covariances with error term} = \begin{bmatrix} \sigma_{0\varepsilon} \\ \sigma_{1\varepsilon} \\ \sigma_{2} \end{bmatrix} = \begin{bmatrix} 0 \\ \sigma_{1\varepsilon} \\ 0 \end{bmatrix}$$

Rewrite (C-4)

$$\Sigma_{xx}\ \text{Bias} = \text{plim}\left[\frac{X'\varepsilon}{T}\right]$$

$$\begin{bmatrix} 0 & 0 & 0 \\ 0 & \sigma_{11} & \sigma_{12} \\ 0 & \sigma_{21} & \sigma_{22} \end{bmatrix} \begin{bmatrix} \text{Bias}(0) \\ \text{Bias}(1) \\ \text{Bias}(2) \end{bmatrix} = \begin{bmatrix} 0 \\ \sigma_{1\varepsilon} \\ 0 \end{bmatrix}$$

or:

$$\sigma_{11}\ \text{Bias}(1) + \sigma_{12}\ \text{Bias}(2) = \sigma_{1\varepsilon}$$

$$\sigma_{21}\text{Bias}(1) = -\sigma_{22}\text{Bias}(2).$$

The second equation can be written as:

$$\text{Bias}(1) = -\frac{\sigma_{22}}{\sigma_{21}}\ \text{Bias}(2).$$

Thus

$$\sigma_{11}\left(-\frac{\sigma_{22}}{\sigma_{21}}\,\text{Bias}(2)\right) + \sigma_{12}\text{Bias}(2) = \sigma_{1\varepsilon}\ ,$$

or

$$\text{Bias}(2) = \frac{\sigma_{1\varepsilon}}{\left(\sigma_{12} - \sigma_{11}\,\dfrac{\sigma_{22}}{\sigma_{12}}\right)}\ .$$

Since

$$\sigma_1 = \text{cov}\left(\varepsilon, \frac{-\varepsilon}{a-1}\right) = \sigma_{\varepsilon\varepsilon}\left[\frac{-1}{a-1}\right]\ ,$$

$$\text{Bias}(2) = \frac{\sigma_{\varepsilon\varepsilon}\left[\dfrac{-1}{a-1}\right]}{\left[\sigma_{12} - \dfrac{\sigma_{11}}{\sigma_{12}}\,\sigma_{22}\right]}\ . \qquad \text{(C-5)}$$

The numerator of Bias(2) as expressed in equation (C-5) is positive since $a < 0$. The denominator is negative since $R^2 \leq 1$ implies $\sigma_{12}^2 \leq \sigma_{11}\sigma_{22}$. Thus, Bias(2) is negative.

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